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THE BEHAVIOUR OF ATHENS STOCK PRICES

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The behaviour of Athens stock prices

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I. INTRODUCTION

In the last two decades a vast amount of research has been conducted on price movements in capital markets. A problem of continuing interest in financial economics analysis is that of evaluating the efficient capital market hypothesis. Briefly stated, the hypothesis claims that a 'market in which prices fully reflect available information is called efficient'; see Fama (1970, An efficient market is one in which the price always incorporates all the information available to the market. This implies that there are no opportunities to make extraordinary profits by exploiting information contained in past price changes. A large number of empirical studies have been developed supporting this hypothesis. Jensen *et al.* (1978) calls it the best established empirical fact in economics.

Numerous investigators have examined the efficient capital market hypothesis in developed countries. There have been few studies on the efficient capital market hypothesis for developing countries.

Hong (1978a, b) tested for weak-form efficiency on the Stock Exchange of Singapore (SES) and he concluded that the random walk hypothesis cannot be rejected for the SES. In addition, Ang and Pohlman (1978) found that the SES is efficient in the weak sense.

Gandhi *et al.* (1980), using a number of well known empirical tests, showed the inefficiency of the Kuwaiti stock market.

All these studies used conventional statistical methodologies which assume normal distribution; however, none of them investigated the nature of the distribution of monthly returns. Therefore, these studies have generated additional interest as to

(a) additional evidence for other developing countries, and
(b) use of proper tests on the form of the distributions of stock returns for investigating the form of stock returns.

The purpose of this paper is to investigate the capital market efficiency in the case of the Athens stock market and to determine to what extent the empirical distribution of monthly returns conforms to the normal distribution. The remainder of this paper is organized as follows. Section II will briefly discuss the methodology and data used. Section III summarizes and reports the results of three tests (autocorrelation coefficients, Ljung–Box statistic and Kolmogorov–Smirnov statistic) employed to test the independence of successive stock price changes in the Athens stock market. Section IV presents and discusses the empirical results of runs analysis and the rank version of von Neumann's ratio test for randomness. The purpose of Section V is to examine the distribution of stock returns in the

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context of Paretian stable distributions (Mandelbrot–Fama hypothesis). Some conclusions are noted in Section VI.

II. THEORETICAL FRAMEWORK AND DATA

The model to be tested in this paper can be formulated as follows. Let $\log p_t$, $\log p_{t+1}, \ldots, \log p_{t+k}$ be successive log prices, $\Phi_t, \Phi_{t+1}, \ldots, \Phi_{t+k}$ be successive information sets, and

$$\left.\begin{array}{l} x_t = \log p_t - E(\log p_t | \Phi_{t-1}) \\ \vdots \qquad\qquad \vdots \\ x_{t+k} = \log p_{t+k} - E(\log p_{t+k} | \Phi_{t+k-1}) \end{array}\right\} \tag{1}$$

be successive returns, and let $E(.|\Phi_t)$ denote the objective expectation conditional on Φ_t. Lower-case letters will denote observed data (p_t, x_t), which are interpreted as realizations of random variables denoted by the corresponding capital letters (P_t, X_t). The information sets, Φ, considered here are the sets of present and past stock prices recorded monthly. The sequence $x_t, \ldots, x_{t+k}$ is a fair game with respect to information Φ if

$$E(x_{t+k}|x_t) = E(x_{t+k}) = 0, \qquad \forall t, k \tag{2}$$

Equation 2 holds if the conditional expected rates of return $E(x_{t+k}|x_t)$ are unbiased in each time period and if individual returns are serially independent.

According to Fama (1970), the 'weak' fair game includes in Φ the information from only the sequence of past values. One implication of this definition is that

$$E(\log p_{t+1}|\Phi_t) = \log p_t \qquad \forall t \tag{3}$$

Therefore Equation 1 becomes

$$\left.\begin{array}{l} x_t = \log p_t - \log p_{t-1} \\ \vdots \qquad\qquad \vdots \\ x_{t+k} = \log p_{t+k} - \log p_{t+k-1} \end{array}\right\} \tag{4}$$

The efficiency criterion, Equation 2, requires that

$$E[(x_{t+k} - x_{t+k-1})|(x_t - x_{t-1})] = E(x_{t+k} - x_{t+k-1}) = 0 \qquad \forall t, k \tag{5}$$

Observe that Equation 5 implies that successive rates of return follow a martingale. In addition, Equation 5 implies that x_t should be uncorrelated with any past information in Φ_{t-1}. Thus the 'weak' efficiency hypothesis will be tested in this paper using (a) the sample autocorrelation function, (b) the 'portmanteau' statistic, and (c) the Kolmogorov–Smirnov statistic; and in order to test for randomness we will examine (d) the runs analysis and (e) the rank version of the von Neumann ratio test. In addition, the paper will examine if the distribution of stock price changes belongs to the non-normal class of stable distributions (Mandelbrot–Fama hypothesis).

The data of this study consist of the monthly closing stock prices for ten companies with shares quoted on the Athens stock market over the period January 1965 to December 1984. The selection criterion was the level of trading activity in the period of consideration.

All the companies except one were included in the construction of the price indices of the Athens stock market. The data were adjusted for stock splits, cash dividends and bonus

issues. These adjustments were carried out in the standard fashion (Fama, 1965a).

In this paper we will be mainly concerned with differences in the logarithm of prices (see Equation 4), or

$$x_t = \log p_t - \log p_{t-1} \tag{6}$$

This is a procedure widely used in most empirical studies for the following reasons:

(a) that x_t represents the yield, with continuous compounding from holding the stock during that month;
(b) it has been shown by Moore (1964) that the variability of simple price changes for a given stock is an increasing function of the price level of the stock; taking logs seems to neutralize most of this price effect; and
(c) it may be remarked that for the runs tests it does not matter whether $p_t - p_{t-1}$ or $\log p_t - \log p_{t-1}$ is used, since only signs, not magnitudes, are involved.

III. TEST OF INDEPENDENCE

Serial correlation analysis

Independence was first investigated by Kendall (1953). Using serial correlation coefficients for the first difference of weekly London stock price indices, Kendall found the estimated coefficients to be very close to zero. He arrived at the conclusion that 'the series looks like a "wandering" one, almost as if once a week the Demon of Chance drew a random number from a symmetrical population of fixed dispersion and added it to the current price to determine next week's price'.

The Kendall method was extended to weekly New York Stock Exchange indices by Cootner (1962) and Moore (1964) and they found some evidence of correlation.

Fama (1965a), using the 30 US companies which make up the Dow Jones industrial index, produced evidence of dependence in the price changes.

Kemp and Reid (1971), using British share price data, applied only non-parametric tests and found some evidence of dependence in the price changes.

Conrad and Jüttner (1973) applied parametric and non-parametric tests to daily stock price changes in the German stock market. They found that the random walk hypothesis is inappropriate to explain the price changes.

The serial correlation technique was employed by Dryden (1970), Jennergren and Korsvold (1974, 1975) and Roux and Gilbertson (1978) in their studies of the weak form efficient market hypothesis.

Similar tests were also carried out in a number of other countries: Praetz (1969, 1973, 1979) examined the behaviour of Australian share prices; Solnik (1973) studied 234 stocks from Belgian, British, Dutch, French, Italian, German, Swiss and Swedish stock markets; Ang and Pohlman (1978) studied 54 stocks from the stock materials of Australia, Hong Kong, Japan, Malaysia and Singapore; and Gandhi *et al.* (1980) analysed the Kuwaiti stock market.

Cooper (1982) studied world stock markets using monthly, weekly and daily data for 36 countries. He examined the validity of the random walk hypothesis by employing correlation analysis, run analysis and spectral analysis. With respect to the USA and the UK, the evidence supports the random walk hypothesis. For all other stock markets, the evidence is less clear.

Recently, Wong and Kwong (1984) examined the behaviour of the daily closing prices of 28 Hong Kong stocks. The results of serial correlation coefficients showed that the successive stock price changes were dependent random variables. The authors concluded that the Hong Kong market is not efficient in the weak form.

Thus, this brief review of the capital efficient market hypothesis indicates some mixed findings regarding the weak form of efficiency of the stock markets.

In order to shed some light on this apparent controversy, in this paper we have chosen a developing country as the basis of our study.

Numerous tests for establishing statistical independence in a stock-price time series are available in the literature. The independence of successive log price changes for the ten stocks has been investigated by means of the sample autocorrelation function (SACF). The SACF, r_k, measures the amount of linear dependence between observations in a time series that are separated by lag k, and is defined as

$$r_k = \frac{\sum_{t=1}^{n-k} (x_t - \bar{x})(x_{t+k} - \bar{x})}{\sum_{t=1}^{n} (x_t - \bar{x})^2} \tag{7}$$

where r_k is the autocorrelation coefficient for a lag of k time units.

If a set of data is independently distributed, the r_k are zero for all time lags of the differenced series.

An approximate formula for the standard error of r_k, $\mathrm{SE}(r_k)$, has been derived by Bartlett (see Kendall and Stuart (1961)):

$$\mathrm{SE}(r_k) = \frac{1}{n^{1/2}} \tag{8}$$

The SACF up to the first ten lags is estimated for each stock. The results of the SACF test are shown in Table 1. The autocorrelations were all quite low, and only a few first order autocorrelations were statistically significant. For the entire sample, 11 SACFs were significantly different from zero at the 5% level. From these 11 SACFs, seven values were significantly different from zero for the first lag. However, this result is not unexpected. As Working (1960) has shown, the monthly averaging of daily random increments will often produce a first-period autocorrelation of 0.25.

Table 1 shows that, after the first lag, the remaining autocorrelations were all small and statistically insignificant and we can conclude that past returns were of no use in predicting the expected value of the next month's returns.

Therefore, the results based on SCAFs provide support for the existence of independence. In order to provide more evidence about the independence of successive log stock-price changes, we employed the Ljung–Box statistic (Ljung and Box, 1978): rather than considering each SACF individually, this is the 'portmanteau' statistic, defined as

$$Q(k) = n(n+2) \sum_{m=1}^{k} \frac{1}{n-m} r_m^2 \tag{9}$$

Under $H_0: r_1 = \ldots = r_k = 0$, Q is asymptotically x^2 distributed with k degrees of freedom. If H_0 is false the test statistic tends to become large, thus indicating model inadequacy.

The last column of Table 1 contains the Ljung–Box statistic for an overall test of uncorrelated data.

Table 1. *Monthly serial correlations for lags 1, 2, . . . , 10*

Stock	Lag 1	Lag 2	Lag 3	Lag 4	Lag 5	Lag 6	Lag 7	Lag 8	Lag 9	Lag 10	Ljung–Box Q
National Bank	0.07317*	0.04441	−0.05442	0.05368	0.04707	0.04133	0.02657	−0.05883	0.03729	0.01938	5.564
Bank of Greece	0.01528	0.01796	−0.01858	0.01147	0.01877	0.05898	0.01723	0.03805	−0.04781	0.2676	2.356
Commercial Bank	0.9246*	0.0156	0.01667	0.05978	0.03377	0.01030	0.00969	0.01903	0.03167	0.02463	3.911
Ionian Popular Bank	−0.02396	0.03045	0.01722	−0.03716	0.05709	−0.06164	0.01154	0.03342	0.01878	0.03618	3.398
Mortgage Bank	0.01045	−0.13512*	0.02923	0.05241	0.05361	−0.02336	−0.03168	−0.01539	0.05818	0.015621	7.799
Credit Bank	−0.07549*	0.05240	−0.03314	0.00737	0.03165	−0.06120	−0.04982	0.01918	−0.00588	0.01226	4.268
Piraiki–Patraiki	0.12677*	0.12765*	−0.03001	−0.0059	0.02795	0.05194	0.00277	0.01294	0.01096	0.02466	9.515
Chemical Fertilizers	0.14580*	0.10964*	0.01492	0.01105	0.01852	0.01634	0.04864	0.02287	0.05109	−0.02363	10.932
General Cement	−0.21855*	0.03188	0.05686	0.01679	0.01534	0.04586	−0.02088	0.01715	0.05173	−0.02297	14.366
Titan (Cement Co.)	0.13652*	−0.05933	−0.07732*	0.03503	0.01409	0.02964	−0.05848	−0.04745	−0.02561	−0.01313	9.608

*Significant serial correlation at the 5% level.

From an examination of the results, it is apparent that the Q statistic is not significant at the 5% level. For example, in the case of General Cement the computed Q statistic was 14.366, as compared with the critical level of X^2_{10} of 18.31. From the results of the Q statistic one could arrive at the conclusion that the overall successive changes in stock prices are independent. This observation suggests that the Athens stock market is weakly efficient.

Kolmogorov–Smirnov test

The Kolmogorov–Smirnov statistic provides an alternative way for white noise processes. The statistic is taken from the cumulated periodogram $p_1, \ldots, p_k$ defined by the time series $X_1, \ldots, X_n$, where $k=n/2$. The cumulated periodogram is given as

$$S_j = \frac{\sum_{h=1}^{j} p_h}{\sum_{h=1}^{k} p_h}, \qquad j=1, 2, \ldots, k \tag{10}$$

The test for autocorrelation suggested by Durbin (1967), which uses the cumulated periodogram (Equation 10), is

$$D_n = \max \left| S_j - \frac{j-1}{k-1} \right| \tag{11}$$

This maximum value is compared with the critical value to determine whether the time series elements $X_1, \ldots, X_n$ are uncorrelated.

The last column of Table 2 shows the Kolmogorov–Smirnov statistics. The critical value at the 0.05 level is approximately 0.191. The statistics in Table 2 indicate that the successive changes in stock prices are independent at the 0.05 level of confidence. Consequently, the Kolmogorov–Smirnov test confirms the results from the Ljung–Box test. Therefore, judging from these results alone, the empirical evidence supports the hypothesis of weak efficiency in the Athens stock market.

Table 2. *Empirical distributions*

Stock	Random test		Normality test				
	RVN	von Neumann	$\sqrt{\beta_1}$	β_2	SR	Skewness[a]	Kolmogorov–Smirnov statistic[b]
National Bank	0.38	0.037	1.9	23.01	12.9	LS	0.120
Bank of Greece	0.077	0.045	2.2	3.33	8.2	RS	0.081
Commercial Bank	0.39	0.096	5.6	66.5	14.2	LS	0.113
Ionian Popular Bank	6.5	0.092	1.4	10.2	9.9	RS	0.092
Mortgage Bank	1.8	0.052	6.8	81.1	15.0	LS	0.141
Gredit Bank	0.84	0.070	7.2	80.2	14.6	LS	0.125
Piraiki–Patraiki	0.23	0.025	0.42	2.16	7.5	LS	0.131
Chemical Fertilizers	0.69	0.040	1.05	8.4	10.3	LS	0.052
General Cement	1.6	0.070	1.8	30.4	14.3	LS	0.074
Titan (Cement Co.)	0.85	0.071	1.7	11.2	9.4	LS	0.115

[a] RS, right skew; LS, left skew.
[b] Critical value 0.191 at 0.05 level.

IV. TESTS FOR RANDOMNESS

Run analysis

The efficient market hypothesis (i.e. that price changes are unsystematic) can be tested by using runs tests. The issue behind every runs analysis in testing randomness is that too many or too few runs (i.e. sequences of price changes of identical sign) are unlikely if the sample is random. Under the hypothesis of independence the expected number of runs can be estimated as

$$m=\left[N(N+1)-\sum_{i=1}^{3} n_i^2\right]\Big/N \tag{12}$$

where N is the total number of observations (changes in stock prices) and the n_i are the number of price changes of each type, with $i=1, 2, 3$ representing the total number of positive $(+)$, negative $(-)$ and zero (0) stock price changes. The variance of m is

$$\sigma_m^2=\frac{\sum_{i=1}^{3} n_i^2\left[\sum_{i=1}^{3} n_i^2+N(N+1)\right]-2N\sum_{i=1}^{3} n_i^3-N^3}{N^2(N-1)} \tag{13}$$

For large N, the sampling distribution of m is approximately normal. The standardized variable may be determined as

$$Z=\frac{(R+0.5)-m}{\sigma_m} \tag{14}$$

where R is the actual number of runs.

Table 2 shows the results of the runs test. For one out of ten companies the hypothesis of randomness was rejected. Thus, for nine out of ten companies, the empirical evidence tends to accept the hypothesis of (weak) efficiency in the Athens stock market.

The rank version of von Neumann's ratio test

An alternative test procedure, the rank von Neumann ratio test (see Bartels (1982)), will be used to assess the randomness of log price differences.

The rank version of the von Neumann's ratio test procedure can be summarized as follows. Let R_i be the rank of the ith observation in a sequence of T observations; then the rank version of von Neumann's ratio (RVN) is defined by

$$\mathrm{RVN}=\frac{\sum_{i=1}^{T-1}(R_i-R_{i+1})^2}{\sum_{i=1}^{T}(R_i-\bar{R})^2} \tag{15}$$

or the numerator of (15), NM:

$$\mathrm{NM}=\sum_{i=1}^{T-1}(R_i-R_{i+1})^2 \tag{16}$$

is an equivalent test statistic under randomization. Bartels has provided tables of the critical values for the rank version of the von Neumann ratio.

The rank version of the von Neumann's ratio test was applied to our data from the Athens stock market. The results presented in Table 2 indicate that in all cases the test statistic RVN is not significant at the 5% level. Judging from the rank version of the von Neumann's ratio test, the efficiency hypothesis must be accepted. In the case of the Ionian Popular Bank, the runs tests rejected the efficiency hypothesis while the von Neumann's ratio (RVN) appear to accept the efficiency hypothesis. Consequently, the von Neumann's test does not confirm the result from the runs test for the Ionian Popular Bank. Bartels (1982) has advanced a relatively straightforward explanation according to which 'runs tests should be less powerful than a test based on ranks since runs tests completely ignore the magnitudes of the observations'. It must therefore be concluded for all the companies of our sample that successive changes in stock prices are random.

V. DISTRIBUTION OF MONTHLY RETURNS

Tests of normality

In this section we investigate whether or not the empirical distributions of successive log stock price changes conform to the normal distribution.

Consider a sample $X_1, \ldots, X_n$. The coefficient of skewness, β_1, is defined as

$$(\beta_1)^{1/2} = \frac{\mu_3}{\mu_2^{3/2}} \tag{17}$$

where

$$\mu_3 = \frac{1}{n} \sum_{i=1}^{n} (X_i - \bar{X})^3$$

and

$$\mu_2 = \frac{1}{n} \sum_{i=1}^{n} (X_i - \bar{X})^2$$

If $X_1, \ldots, X_n$ constitute a random sample from a normal population, then $(\beta_1)^{1/2}$ is approximately normally distributed, with zero mean and standard error $\mathrm{SE}(\beta_1)^{1/2} = (6/n)^{1/2}$. Consequently, the ratio $(\beta_1)^{1/2}/\mathrm{SE}(\beta_1)^{1/2}$ can be compared with the standard normal variate to test the hypothesis of normality. For a normal distribution N(0, 1), $(\beta_1)^{1/2} = 0$. Geometrically, negative skewness ($\mu_3 < 0$) is seen as an extended tail to the left (left skew (LS)), and positive skewness ($\mu_3 > 0$) implies an extended tail to the right (right skew, (RS)).

The coefficient of kurtosis, β_2, is defined as

$$\beta_2 = \frac{\mu_4}{\mu_2^2} - 3 \tag{18}$$

For large values of n, β_2 is normally distributed, with mean zero and standard error $\mathrm{SE}(\beta_2) = (24/n)^{1/2}$ when $X_1, \ldots, X_n$ are a random sample from a normal population (Kendall and Stuart, 1969). A normal distribution N(0, 1) has $\beta_2 = 0$. A distribution with positive kurtosis has a sharper peak than the normal distribution, whereas one with negative kurtosis is relatively flat.

The skewness and kurtosis coefficients are shown in Table 2. The observed distributions of successive log stock price changes all have kurtosis coefficients considerably larger than 3, a condition known as leptokurtosis. Using the criteria $|(\beta_1)^{1/2}/\mathrm{SE}((\beta_1)^{1/2})|<2$ and $|\beta_2/\mathrm{SE}(\beta_2)|<2$ to conclude in favour of normality, it is seen from Table 2 that all ten stocks display substantial peakedness. It is noteworthy that the empirical distributions of successive log stock price changes in the case of the Athens stock market are not drawn from a normal process.

It can reasonably be concluded that eight of the ten values of μ_3 are negative, implying negative skewness (LS) (see Table 2).

The normality hypothesis has been rejected by several researchers; see, for example, Fama (1965a), Dryden (1970), and Jennergren and Korsvold (1974).

The results reported for $(\beta_1)^{1/2}$ and β_2 (see Table 2) lead to the conclusion that log stock price changes have non-Gaussian distributions. The coefficient of kurtosis for a sample is not a reliable indicator because in the presence of outliers the kurtosis coefficient can rise very quickly. In order to provide further evidence concerning the non-normality hypothesis, we use the Studentized range (SR) test to examine the normality null hypothesis. The SR statistic for the observations $X_1, \ldots, X_n$ is given by

$$\mathrm{SR}=\frac{\max\limits_{1\leqslant i\leqslant n}\{X_i\}-\min\limits_{1\leqslant i\leqslant n}\{X_i\}}{\left[\dfrac{1}{n-1}\sum\limits_{i=1}^{n}(X_i-\bar{X})^2\right]^{1/2}} \tag{19}$$

where $\bar{X}=\dfrac{1}{n}\sum\limits_{i=1}^{n}X_i$.

Tables of significance points for this statistic are given by Pearson and Hartley (1970). In addition, Fama and Roll (1971) have provided evidence indicating that the Studentized range seems to be a good technique for goodness-of-fit tests of normality against non-normal stable alternatives.

Table 2 summarizes the SR statistic for each firm. The general impression one gains from Table 2 is that the Studentized ranges for all monthly log price changes are well beyond the critical values. This result, at least for the sample considered, suggests that monthly distributions of log stock price changes belong to the normal class of stable distributions. Therefore, these results are consistent with a heavy-tailed non-Gaussian distribution.

A natural way is then to investigate if departures from normality are in the direction of Mandelbrot's (1963, 1967) hypothesis of stable Paretian returns. Therefore, continuing effort must be made to examine if the Paretian stable distribution appears to be appropriate when applied to stock price data for a developing country.

The Mandelbrot–Fama hypothesis

In his classic paper on 'Brownian motion in the stock market', Osborne (1959) was the first investigator of the probability distribution of stock price changes. Assuming that price changes across transactions are identically and independently distributed (i.i.d.) random variables with finite variance, and that transactions occur in large numbers over a period of, say, months, then the monthly price changes will be the sum of i.i.d. random variables and, by the central limit theorem, they must converge to a normal distribution.

A more serious challenge to the normality assumption has been voiced by Fama (1965a). He has shown that the empirical distribution of daily price changes for 30 New York Stock

Exchange securities is leptokurtic. Fama's findings suggest that there is a discrepancy between the empirical distribution and the normality assumption.

Since his work, several researchers have tried to explain this discrepancy. Mandelbrot (1963) observed that the mean square value of the difference in the log of cotton prices never approached a limiting value. According to Mandelbrot, this information implies that a member of the family of stable Paretian distributions would serve as the probability model for the distribution of the changes in the log of prices.

The properties of the stable distribution have been extensively investigated: see Levy (1940) and Gnedenko and Kolmogorov (1968). The interest in the stable distributions stems from the fact that the sum of n independent random variables with probability density functions $f_1(t), \ldots, f_n(t)$ has a density function $f_s(t)$ given by the n-fold convolution

$$f_s(t) = f_1(t) * f_2(t) * \ldots * f_n(t)$$

where $*$ denotes the convolution operation.

Khintchine and Levy (1936) have advanced the following theorem: a necessary and sufficient condition for a distribution function $F(x)$ to be stable is that the logarithm of its characteristic function $\Phi(t)$ can be represented by

$$\log \Phi(t) = \mathrm{i}\delta t - \gamma |t|^{\alpha}[1 + \mathrm{i}\beta(t/|t|)\omega(t, \alpha)] \tag{20}$$

where

$$\omega(t, \alpha) = \begin{cases} \tan \dfrac{\alpha\pi}{2} & \text{if } \alpha \neq 1 \\ \dfrac{2 \log |t|}{\pi}, & \text{if } \alpha = 1 \end{cases}$$

and $0 < \alpha \leqslant 2$, $-1 \leqslant \beta \leqslant 1$, $\gamma \geqslant 0$, δ is any real number and the characteristic function $\Phi(t)$ is the Fourier transform of the distribution function $F(x)$.

The parameters δ, γ, α and β indicate the location, dispersion, tailedness and skewness of the distribution respectively. The parameter α is called characteristic exponent of the distribution and determines the type of stable distribution. The normal distribution ($\alpha = 2$) and the Cauchy ($\alpha = 1$) are members of the stable family. If $\beta = 0$ the distributions are symmetric about the location parameter. When $\beta > 0$ the distribution is skewed right, and when $\beta < 0$ it is skewed left.

Mandelbrot (1963) and Fama (1963, 1965a, 1965b) hypothesized that the distribution of log price changes is best fitted by the stable laws. Therefore, according to the Mandelbrot–Fama hypothesis, the distribution of returns is stable Paretian with characteristic exponent $1 < \alpha < 2$. This hypothesis suggests that the stock return distributions have 'infinite' theoretical variance and that the distribution shapes are stable.

Most of the available empirical evidence on the characteristic exponent refers to the developed economies. In this section we report additional empirical evidence based on data from a developing country.

Fama and Roll (1968, 1971) have provided methods for estimating the parameters of stable distributions. In addition, they have generated probability tables of the symmetric stable distributions. Fama (1965a) estimated α and found it to be less than 2; thus he concluded that the distribution of monthly returns belonged to a non-normal member of the stable family of distributions. Next we turn to the problem of estimating the parameters of the stable distribution. The estimation procedure for the crucial parameter α which we used

was suggested by Fama and Roll (1971). The estimate of the characteristic exponent, α, can be obtained using the ratio

$$Z_f = \frac{X_f - X_{1-f}}{X_{0.72} - X_{0.28}} \times 0.827 \tag{22}$$

where X_f is of $(.f)(n+1)$th order statistic. Fama and Roll have suggested that $f=0.99$ serves well to determine whether or not $\alpha=2$, and that $f=0.95-0.97$ is robust in estimating α.

In this paper the f fractile of 0.96 is used because values of $0.95 \leqslant f \leqslant 0.97$ have minimum dispersion. Thus $\hat{Z}_{0.96}$ can be used to estimate $\hat{\alpha}$. This value of $\hat{Z}_{0.96}$ is then referred to a table of standardized symmetric stable cumulative distribution functions to obtain the estimate of α whose 0.96 fractile best matches $\hat{Z}_{0.96}$.

Table 3 gives the results of estimating α for fractiles $f=0.96$ and $f=0.99$. An inspection of Table 3 indicates that α varies between firms, and that all 'tail' exponents lie between 1 and 2. Fama and Roll (1971) report the results of a Monte-Carlo experiment to test hypothesis $H_0 - \alpha = 2$ – against hypothesis $H_1 - \alpha < 2$: the results show that the null hypothesis is rejected at the 0.05 level in 17 cases out of 20. Table 3 shows that approximately 85% of the 20 estimates of characteristic exponents are less than 1.7. As indicated by Fama (1965a), if α is less than 1.7 then the data provide support for the existence of independence. The findings of the estimates of the characteristic exponent, α, suggest that distributions of monthly log stock price changes belong to the non-normal class of stable distributions. Hence in the case of the Athens stock market, the stock return distributions are consistent with the Mandelbrot–Fama hypothesis.

VI. CONCLUSIONS

In this paper some results have been presented on the weakly efficient capital market hypothesis in the case of the Athens stock market. In a test of the null hypothesis that the observations are random against the alternative that they are not random, the null hypothesis is accepted using both runs tests and the rank version of the von Neumann test.

Table 3. *Estimates of characteristic exponents*

Stock	$\hat{\alpha}_{0.96}$	$\hat{\alpha}_{0.99}$
National Bank	1.20*	1.51*
Bank of Greece	1.22*	1.63*
Commercial Bank	1.05*	1.55*
Ionian Popular Bank	1.01*	1.12*
Mortgage Bank	1.14*	1.63*
Credit Bank	1.30*	1.52*
Piraiki–Patraiki	1.58*	1.70
Chemical Fertilizers	1.79	1.83
General Cement	1.40*	1.65*
Titan (Cement Co.)	1.35*	1.56*
Average $\hat{\alpha}$	1.304*	1.57*

* $\hat{\alpha}$ significantly less than 2 at the 0.05 level.

The results based on sample serial correlation (SACF) to examine independence are weak. In order to provide more evidence about independence, the Q statistic and the Kolmogorov–Smirnov statistic have been used. According to the Q statistic there is evidence that the stock returns in the Athens stock market are independent. These results then are consistent with the weakly efficient market hypothesis. In addition, this conclusion is supported by the results of the Kolmogorov–Smirnov statistic. Therefore, the overall evidence tends to support the weak form of the efficient market model. The predominant pattern in Table 3 is that the stock return distributions are consistent with a stable Paretian distribution hypothesis: that is, the data appear to support the Mandelbrot–Fama hypothesis. Moreover, examination of the characteristic exponents indicates no underlying characteristic exponent common to all stocks. This also implies that existing portfolio models based on non-normal stable distributions are inappropriate (e.g. Fama 1965b, 1971).

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